How Coalitions Shape Legislative Institutions in Parliamentary Democracies

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Abstract

Previous work shows that robust legislative oversight institutions strengthen the ability of coalition governments to enforce policy agreements. This raises the question of whether coalitions choose such institutions strategically. In this article, I propose a simple formal model of procedural choice under coalition government and probe its empirical plausibility in a novel analysis of committee procedure reforms undertaken in 14 European parliamentary democracies over more than 60 years. The analysis reveals an empirical pattern which is broadly consistent with expectations. Procedural opportunities to use legislative committees to challenge ministerial proposals tend to be strengthened under multi-party cabinets in which preference divergence between coalition partners is substantial and in which parties have similar probability of making proposals to implement joint policies. These findings open interesting avenues for future work on how parties shape legislative institutions in parliamentary democracies.

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Recent research shows that legislative institutions influence the ability of parliamentary coalition governments to enforce policy compromises (Kim and Loewenberg 2005; Martin and Vanberg 2011, 2020; Carroll and Cox 2012; Fortunato, Martin, and Vanberg 2017). A key finding in this area is that enforcement of coalition commitments is facilitated by legislative procedures that enable parties to challenge and revise ministerial proposals and to acquire independent policy-relevant expertise. Most of this work treats legislative institutions as an exogenous environment. This is a useful simplifying assumption but scholars are well aware that legislative rules are not set in stone (Diermeier and Krehbiel 2003; Krehbiel 2004; Shepsle 2017). In fact, procedures in many parliamentary democracies can be changed with a simple majority (Sieberer, Müller, and Heller 2011). Legislative institutions are thus self-imposed constraints chosen by majorities whose behaviour they are meant to motivate and guide (Calvert 1995; Diermeier, Prato, and Vlaicu 2020). This raises the issue of whether parties participating in governing coalitions adjust rules strategically to reflect their needs.

Studying this question is an important task for at least three broad reasons. First, the creation of institutions to address tensions that arise in multi-party governments is of core interest in coalition research (Thies 2001; Müller and Meyer 2010; Strøm, Müller, and Smith 2010). By examining the design of legislative procedures, I thus contribute to the voluminous literature on the mechanisms of coalition governance (Lipsmeyer and Pierce 2011; Indridason and Kristinsson 2013; Klüver and Bäck 2019; Höhmann and Sieberer 2020). Second, endogenous procedural choice is an issue of clear and continuing interest to political scientists (Binder 1996; Schickler 2001; Krehbiel 2004; Diermeier and Vlaicu 2011; Goet 2021; Sieberer et al. 2020). Studying how multi-party governments shape legislative institutions means I can shed light on institutional development in settings in which parties govern

together but contest elections separately. Third, understanding the choice of legislative procedures allows me to contribute to the broader debate about the evolution of critical political institutions including constitutions (Elkins, Ginsburg, and Melton 2009; Tsebelis 2022) and electoral systems (Boix 2004; Benoit 2004; Cusack, Iversen, and Soskice 2007). By analyzing the conditions under which parties choose to construct strong oversight institutions, we can gain important insights into how consensual constitutional designs emerge in parliamentary democracies (Powell 2000; Lijphart 2012).

My specific concern is the choice of committee procedures under coalition governments. The strength of committee scrutiny powers is a key institutional feature that is found to moderate the ability of coalitions to enforce policy deals (Martin and Vanberg 2011, 2020). Previous work suggests that because they benefit from committee strength, parties that frequently participate in multi-party governments may be motivated to establish and maintain strong committees. There is indeed some evidence that committee strength is associated with the prevalence of multi-party government (Powell 2000). In a comparison of 15 western European democracies, Hallerberg (2000) shows that countries in which multi-party coalitions are the norm tend to have procedures that facilitate committee monitoring of ministerial activities; in contrast, such procedures are largely absent in countries with single-party majority governments. In a recent study of 31 countries, André, Martin and Depauw (2016) demonstrate that countries that experienced a high rate of coalition government in 1979-2009 tended to have stronger legislative committees at the end of that period. These authors also complement their analysis with supportive evidence from a case study of the Irish committee system (see also Martin and Vanberg 2011, 158–9).

Beyond such broad patterns, however, we know relatively little about the conditions under

which coalitions implement procedural reforms to facilitate committee oversight. Theorizing and empirically investigating this dynamic is thus an important next step in the study of endogenous committee procedures in parliamentary democracies. Previous work offers some interesting pointers in this regard. Martin and Vanberg (2011) emphasize the importance of decision-making costs in shaping the conditions for committee procedure reforms. In particular, these authors suggest that procedures can be difficult to change in political systems in which legislative institutions are enshrined in the constitution. Wang (2013) argues that while all coalition cabinets have strong incentives to strengthen legislative committees, the benefits from such a strengthening are particularly attractive to coalition parties that are divided on many policy issues and are more fractionalized and undominated by a single large party. Zubek (2015) argues that governments whose supporting parties are ideologically divided are especially motivated to seek expansions of committee scrutiny powers; in contrast, ideologically unified cabinets are likely to avoid strengthening legislative committees to minimize delays to government business.

Building on such previous insights, I introduce in this article a simple formal model of procedural change under multi-party government. Its main theoretical insight is that coalitions can be expected to expand committee scrutiny procedures more extensively when the preference divergence between the governing parties is large, and that this effect is stronger when parties have more similar probability to make proposals to enforce coalition policy. To probe the plausibility of these theoretical claims, I undertake a systematic analysis of the reforms of committee procedures carried out under multi-party governments in 14 western and east-central European democracies since 1945 or date of democratization. This analysis reveals an empirical pattern which is broadly consistent with my expectations.

The article proceeds as follows. The next section presents a game-theoretic model of procedural choice and derives the main theoretical propositions. The third section contains the empirical analysis, discussing in turn the modelling approach, data and operationalization, and the main results. The concluding section considers avenues for further work on how parties choose procedures in parliamentary democracies.

THEORY

Consider a simple game with two players – M (Minister) and P (Partner) – whose joint consent is required to implement a previously agreed policy compromise. The game proceeds as follows. M proposes a bill which reflects either their ideal policy or the agreed coalition policy. P cannot observe whether the bill proposed by M reflects M's ideal policy or the agreed coalition policy, but P can scrutinize the bill at some cost. If P decides to scrutinize and finds that the bill reflects M's ideal policy, the bill is amended to the agreed policy at some cost to M. If P decides not to scrutinize and the bill proposed by M reflects M's ideal policy, then M's ideal policy is adopted.[1] This simple game is a variant of an enforcement game popular in coalition research (Martin and Vanberg 2011; see also Martin and Vanberg 2005; Franchino and Wratil 2019).[2] A game of this type captures well an important tension confronting multi-party governments: parties cannot observe whether proposals comply with agreed coalition policy when bills originate outside their portfolio assignment.

I assume that players are motivated by the policy which is adopted, position-taking

1. To keep things simple I assume a binary strategy space; in Appendix C, I show that an analogous game with a continuous strategy space leads to similar conclusions.

2. See McCarty and Meirowitz (2007, 177–179) for a general presentation.

incentives, and the costs that players incur during policy-making. Player M receives a policy payoff which is x if a bill reflecting M's ideal policy is adopted, and 0 otherwise. If M's original proposal reflects their ideal policy but the bill is subsequently amended to the compromise policy, M incurs a penalty $k > 0$ which captures the opportunity (and possibly audience) costs incurred in developing a proposal which is ultimately revised. While reneging on the policy compromise entails a cost, by deviating M also gains a position-taking benefit. Parties have incentives to signal to voters that they are staying true to the policies announced at election time. By proposing a bill which reflects their ideal policy, M can make such an electorally beneficial statement.[3] I define this position-taking benefit as θx where $0 < \theta < 1$ captures the importance of position-taking for the minister, and I assume that $k > \theta x$ implying that proposing their ideal policy is not a dominant strategy for player M.[4]

Player P receives a policy payoff which is $-x$ if a bill reflecting M's ideal policy is adopted, and 0 otherwise. The preference divergence between the two players thus increases with the parameter x.[5] If P scrutinizes M's proposal, P pays a cost which has two components. The

3. Much research shows that ministers in multi-party democracies can electorally benefit from such opportunistic non-compliance with previously agreed compromises (see e.g. Martin and Vanberg (2005, 2011)). In the model this position-taking benefit accrues to M during the legislative process.

4. In Appendix D, I show that a game in which $k < \theta x$ leads to similar conclusions.

5. A model with symmetric policy preferences is consistent with the equal-influence model of coalition government policy (see e.g. Tsebelis 2002). In Appendix E, I show that my theoretical results continue to hold under the alternative proportional-influence model (see e.g. Indridason 2011) with reasonable assumptions.

first component captures how costly scrutiny is for the partner under the existing legislative procedure. The procedure shapes the scrutiny cost because it determines what actions the partner is permitted, mandated or prohibited to take during oversight. The procedure thus imposes constraints on the extent of control that the partner has over the scrutiny process.[6] I let ρ be the legislative procedure such that $0 < \rho < 1$ and that when ρ increases the scrutiny cost for the partner increases.[7] The second component captures how costly scrutiny is for the partner given that by engaging in oversight the partner incurs a position-taking cost. Like M, player P faces electoral incentives to stay true to its ideal policy. However, by insisting on the coalition compromise as their preferred benchmark, P is forced to take a position which deviates from that announced at election time. Scrutiny can thus be electorally costly for P because they risk being viewed as 'fighting for the wrong thing', and the size of this penalty increases the more distant the policy compromise is from P's ideal point.[8] I thus define

6. There is much previous work which stresses the importance of legislative procedures for shaping the cost of scrutiny in multi-party democracies (see e.g. Carroll and Cox (2012) and Martin and Vanberg (2020)).

7. The procedural cost is assumed to be a linear function of the legislative procedure ρ. I present results for a more general cost function $f(\rho)$ in Appendix F.

8. The assumption that compromise enforcement has a negative effect for P is consistent with previous work. First, scholars find evidence that voters dislike compromises and punish compromising parties (Fortunato 2017, 2021). These adverse effects arise because parties that compromise (or insist on compromises) are perceived as having shifted their stance which carries a penalty for low representational competence. Second, prior work also shows that legislative scrutiny has electoral consequences. While research has stressed chiefly the opportunities that scrutiny brings for beneficial differentiation (Fortunato 2019), legislative

P's position-taking cost as being equal to x.[9] Combining both components, I define the scrutiny cost as ρx, where ρ is the legislative procedure and x is the preference divergence. The multiplicative functional form indicates that parameters ρ and x are complements in the total oversight cost function. This is a reasonable assumption given that more efficient procedures which give the partner more control over scrutiny ensure a smoother oversight process which in turns lowers electoral costs to the partner.[10]

This game has no pure strategy Nash equilibria. The mixed strategy Nash equilibrium is that M proposes their ideal policy with probability ρ and the coalition compromise with probability $1-\rho$, while P scrutinizes with probability $\frac{x+\theta x}{x+k}$ and accepts the bill without

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scrutiny can also have negative electoral effects (Martin and Vanberg 2011, 54). In my model I focus on the latter, but as I show in Appendix F my results continue to hold when I include positive differentiation benefits.

9. The game is based on the equal-influence model of coalition policy, and both M and P have ideal points which are at the same distance x from the compromise. In Appendix E, I analyze a coalition game based on the proportional-influence model showing that my theoretical results continue to hold with reasonable assumptions. I acknowledge that both these standard models assume that the position-taking cost for P increases as M becomes more extreme. Whether my theoretical results hold under some other alternative models of coalition government policy is an open question, and I leave this issue for future investigation.

10. The assumption that P's scrutiny cost is multiplicative in ρ and x is important for deriving my theoretical results. Assuming, for example, that P's cost is a function of ρ instead of ρx, and that $\rho > x$, results in a pure strategy equilibrium in which M always adopts their ideal policy and P never scrutinizes.

scrutiny with probability $1 - \frac{x+\theta x}{x+k}$.[11] See Appendix A. This equilibrium result accords well with intuition. The Minister proposes their ideal policy with probability which increases as the oversight procedure become more costly (ρ goes up). The Partner engages in scrutiny with probability which increases as preference divergence x and position-taking importance θ grow but declines as the penalty k becomes larger.

Consider now a coalition of two parties a and b that expect to play the policy enforcement game as described above. Assume that party a is expected to act as M with probability β and party b with probability $1 - \beta$. Each party's expected payoff from playing the enforcement game is a function of the legislative procedure, the policy payoff that the party obtains by having its ideal policy implemented and the probability that its ministers are recognized to make policy proposals. Formally, the respective payoffs for parties a and b are $E(u_a) = -(1-\beta)\rho x$ and $E(u_b) = -\beta\rho x$.[12] The coalition parties' expected payoffs are thus a function of the penalty that parties incur when not acting as minister in the enforcement game. Importantly, *both* parties' payoffs strictly *increase* as the legislative procedure becomes *less costly* (i.e. ρ becomes smaller). This latter insight accords well with a theoretical intuition that coalition parties benefit from legislative procedures that facilitate oversight.

Suppose now that parties a and b have control over procedure ρ, and when the enforcement game is to be played, they have an opportunity to reshape the procedural environment. Assume further that when undertaking a procedural reform each party incurs some cost which reflects how well entrenched procedures are in a given political system and the effort required

11. Note that since by assumption $k > \theta x$ then $\frac{x+\theta x}{x+k}$ is a proper probability such that $0 < \frac{x+\theta x}{x+k} < 1$.

12. See Appendix B for the derivation of these equations and those that follow below.

to elaborate new procedural rules. Formally, I define the reform cost as a function of the squared size of the procedural change $(\rho_{sq} - \rho)$ where ρ_{sq} is the status quo procedure and ρ is the newly adopted procedure.[13] The cost of reforms thus rises with the size of the departure from the status quo at an increasing rate, a justified assumption since implementation complexity is likely to be higher for more radical reforms. Given these assumptions, and by the first-order condition, the parties' most preferred new procedures are:

$$\rho_a^* = \rho_{sq} - \tfrac{1}{2}(1 - \beta)x \tag{1}$$

$$\rho_b^* = \rho_{sq} - \tfrac{1}{2}\beta x \tag{2}$$

Equations (1) and (2) reveal that while both parties prefer to lower ρ in equilibrium, they have different preferences regarding the magnitude of this change depending on the probability that they act as minister in the policy enforcement game.[14] As β increases party a prefers smaller deviations while party b prefers larger departures from the existing procedure, and *vice versa*. It is only in the singular case when $\beta = 0.5$ that parties' preferences are identical.

To model how such preference divergences are resolved I propose a simple form of procedural decision-making. Suppose that the agreement of both party a and b is required to amend the status quo procedure and no third party holds a veto over the reform. Suppose further that neither party agrees to undertake a reform if its marginal cost exceeds the

13. Assuming that the reform cost is quadratic in the size of the procedural change simplifies derivation, but my results hold for any function $h(|\rho_{sq} - \rho|)$ which is twice differentiable and is such that $h'(|\rho_{sq} - \rho|) > 0$ and $h''(|\rho_{sq} - \rho|) > 0$ over the whole domain. See Appendix G.

14. Note that parties' preference for lowering ρ is conditioned on a payoff surplus (see Appendix B).

marginal benefit for the party, a reasonable assumption if both parties engage in optimizing behaviour.[15] To the extent that this simple model holds parties may be expected to implement the smaller preferred deviation from the procedural status quo. Intuitively, the party with the larger probability of making proposals will not support a more radical change because it is not worth the extra reform effort. Formally, the procedure ρ^* that parties implement is:

$$\rho^* = \begin{cases} \rho_{sq} - \frac{1}{2}\beta x & \text{if } \beta < 0.5 \\ \rho_{sq} - \frac{1}{4}x & \text{if } \beta = 0.5 \\ \rho_{sq} - \frac{1}{2}(1-\beta)x & \text{if } \beta > 0.5 \end{cases} \tag{3}$$

This result holds two key implications for how coalition governments shape legislative institutions. First, coalition parties are found to be motivated to establish decreasingly costly oversight procedures as the preference divergence between coalition parties increases. This finding resonates well with qualitative arguments in previous research about the impact of intra-coalition ideological divisions on procedural choices. It has been argued that as parties become more divided, potential gains from opportunistic behaviour increase and parties must exert more scrutiny effort; hence, less costly scrutiny procedures benefit both parties (cf. Zubek 2015). Second, the result also implies that the effect of preference divergence is moderated by the probability that coalition parties act as proposers in the coalition enforcement game. The strength of this effect increases as parties become more *similar* in their proposing probabilities (i.e. as $|\beta - (1-\beta)|$ converges to 0). This theoretical result chimes

15. This simple model of procedural decision-making does not allow for parties to make side-deals. I intend to explore such extensions in future work.

well with intuitions expressed in earlier work about the role that coalition fractionalization has in shaping procedural preferences under multi-party coalitions (cf. Wang 2013). Overall, my theory leads me to formulate the following two propositions:

Proposition 1. *Coalition parties adopt decreasingly costly oversight procedures as preference divergence between parties increases.*

Proposition 2. *The effect of preference divergence becomes stronger as coalition parties become more similar in their proposing probabilities.*

DATA AND METHOD

Empirical Task

I examine the plausibility of my theoretical model by analyzing observational data on legislative oversight procedures implemented during the life of parliamentary coalition governments that negotiated joint policy compromises. I reason that if the model captures the realities of institutional choice under coalitions well, I should find two distinct patterns. First, I expect to observe decreasingly costly scrutiny procedures to be implemented as policy preference divergence within coalitions increases. Second, I anticipate that this association is stronger when coalition parties are similar in their proposing probabilities.

A simple statistical test of these expectations involves regressing the legislative procedure implemented in a sample of multi-party governments on the interaction of parties' preference divergence and their proposing probability similarity controlling for constituent interaction terms, the status quo procedure and a set of relevant controls. Using maximum likelihood estimation, and assuming normal distribution for the response, I can obtain my estimates

by maximizing the log-likelihood function given as

$$LL = \sum_{i=1}^{n} \ln \left\{ \frac{1}{\sigma\sqrt{2\pi}} \exp \left(-\frac{(\rho_i - x_i\alpha)^2}{2\sigma^2} \right) \right\} \tag{4}$$

where ρ_i is the legislative procedure which the coalition implements, x_i is a matrix of covariates, and α is a vector of regression coefficients.

An important concern with this simple modelling approach is non-random sample selection. We observe the legislative procedure introduced by the coalition that formed in each government formation opportunity, but procedures are unobserved for all other potential governments. The inferences will be biased if some unobserved variables influence both the government formation and the choice of procedures, and hence modelling procedural choice may require accounting for the government selection process. To address this inferential concern, I follow the approach proposed by Chiba, Martin and Stevenson (2015) and estimate the joint likelihood of government selection and procedural choice using a Gaussian copula.[16] Models based on copula functions have increasingly been used to account for non-random sample selection in political science and beyond (see e.g. Wojtys, Marra, and Radice 2018; Braumoeller et al. 2018; Chiba and Johnson 2019; Ecker and Meyer 2020). In the present

16. Chiba, Martin and Stevenson (2015) estimate a joint model in which the primary dependent variable is government duration. I revise their probability model and the relevant R code to model a continuous response from the normal distribution. See Appendix H for details.

case, the log-likelihood to be maximized to obtain estimates from the joint model is thus

$$LL = \sum_{i=1}^{n} \ln \left\{ \frac{1}{\sigma\sqrt{2\pi}} \exp\left(-\frac{(\rho_i - x_i\alpha)^2}{2\sigma^2} \right) \times \Phi\left(\frac{\Phi^{-1}(u_i) - \Phi^{-1}(v_i)\theta}{\sqrt{1-\theta^2}} \right) \right\} \tag{5}$$

where u and v are univariate cumulative distributions for the selection outcome and the implemented procedure respectively, Φ is the cumulative Gaussian distribution, Φ^{-1} is the inverse cumulative Gaussian distribution, and θ is a correlation coefficient in the bivariate cumulative standard normal distribution.

Legislative Procedures

In measuring legislative institutions I take advantage of a newly published time series of codified rules governing committee procedures in 17 western and east-central European democracies since 1945 or the date of democratisation (Zubek 2021). These data introduce the Committee Policing Strength Index (CPSI) which captures the extent to which committee procedures enable coalition parties to conduct oversight of ministerial proposals. The strength of committee scrutiny powers is commonly considered as an important institutional feature that makes legislative oversight more or less costly for coalition parties (Martin and Vanberg 2011, 2020). By focusing on this type of legislative institutions I am thus able to operationalize the parameter ρ from my theoretical model.

The CPSI index captures committee strength in two dimensions both of which are relevant for my purposes. The first dimension is that of the extent of control that coalition parties have over committee referrals of government bills in the legislative procedure. Formal rules may give parties various levels of influence over committee referrals, ranging from prohibiting referrals to mandating them in all situations. The second dimension is that of scrutiny-

enhancing committee features. The CPSI index relies on a set of procedural features, such as committee specialization or rewrite authority, all of which are commonly associated with how difficult it is for parties to challenge and revise government bills.

The theoretical range of the CPSI index is between 0 and 5, with higher values indicating greater policing strength. For robustness tests, I separate the components of the index and, as suggested in Zubek (2021), I create two partial indices: CPSI-P which captures only the extent of partner control and CPSI-M which measures only scrutiny-enhancing features.

Government Data

The government data for this analysis come from the European Representative Democracy Data Archive (ERDDA) which contains cabinet-level data for European democracies (Hellström, Bergman, and Bäck 2021; Bergman, Ilonszki, and Müller 2019; Bergman, Bäck, and Hellström 2021). I use this source to create an initial dataset of 348 governments that held office in the European countries covered by the CPSI data.[17] From these data I then exclude governments which (i) were non-partisan, (ii) were formed by a single party, or (iii) left office after 31 December 2011 (the cut-off date for the CPSI data). I also exclude coalitions with no explicit policy agreement as my theory does not apply to such governments.[18] I consider a coalition to be based on a policy agreement if its members negotiated compromises ranging from a few selected policies to a comprehensive policy platform. My government

17. I exclude Spain and the UK as these countries do not normally experience coalition government.

18. I present the results for all coalitions regardless of whether ERDDA codes them as having an explicit policy agreement in Appendix M. The main effects remain substantively the same.

dataset thus consists of 174 coalitions in 14 European states. See Appendix I.

I combine the cabinet-level data with the CPSI time series by coding the CPSI estimate for the start and end dates of each government. This creates two key variables for my analysis: 'Status quo CPSI' indicating the procedure inherited from previous government and 'Implemented CPSI' denoting the implemented legislative procedure. To estimate the model specified in Equation (5) I further construct a dataset of all possible governments that could have formed in each of the 174 coalition formation opportunities. I do this by finding all possible combinations of legislative parties that held at least one seat in parliament excluding independent MPs. All together, my final dataset thus contains information about 64,714 potential governments for 174 formation opportunities, each with one government that actually formed. For each of such potential cabinets I obtain a set of various characteristics as described below which will serve as independent variables in the selection equation.

Main Covariates

My main covariate in the models of the implemented procedure is an interaction between the parties' policy preference divergence and their similarity in proposing probabilities. I operationalize the parties' preference divergence using the absolute ideological distance between the positions of the most extreme parties in a coalition. I compute this value using the 2021a release of the Manifesto Project RILE party position data (Volkens et al. 2021) and applying the logit scale proposed by Lowe et al. (2011).

To operationalize the similarity in coalition parties' proposing probabilities, I use legislative seat share as a proxy for a party's probability of acting as proposer in the enforcement game. This choice seems reasonable as legislative strength determines resources such as

office, finance and expertise that shape a party's propensity to introduce policy proposals. It is well known, for example, that coalition parties tend to obtain ministerial portfolios proportionately to their seat shares (Gamson 1961), and portfolio allocation influences how often a party proposes new policies.[19] I obtain the *proposing probability similarity* (PPS) for each government by first computing the mean of seat share differences between the largest party and each coalition party and then normalizing these values to the range (0,1) with larger values indicating higher probability. Formally, the similarity score is computed as

$$PPS = 1 - \frac{\sum_{i=2}^{n} (s_1 - s_i)}{s_1(n-1)} \tag{6}$$

where $\{s_1, s_2, ..., s_n\}$ is an ordered set of coalition parties' seat shares with s_1 denoting the largest party, and n is the number of coalition parties.[20]

For the government selection equation I use a set of ideological, size-related and encumbency covariates which previous research has found as important predictors for coalition formation outcomes (see Laver and Schofield 1990; Martin and Stevenson 2001, 2010). Specifically, I model government selection as being conditioned by *minority government, minimal-winning status, median party in government, the number of government parties, ideological divisions within government, previous PM party in government* and *incumbent*

19. In principle portfolio shares could also be used as a proxy for proposing probability. I present results with PPS based on portfolio shares in Appendix N. See also the discussion under robustness.

20. For robustness I also run my analysis with alternative PPS specifications. See Appendix N for details.

coalition. I construct all of these independent variables using the data available from the ERRDA archive and the Manifesto Project (Hellström, Bergman, and Bäck 2021; Bergman, Ilonszki, and Müller 2019; Bergman, Bäck, and Hellström 2021; Volkens et al. 2021).

Controls

I include a number of control variables when modelling the procedure implemented by coalition governments that actually formed. The first control is an estimate for the status quo committee strength inherited from a previous government. The inclusion of this variable follows directly from my theoretical model as given in Equation (3). Second, I control for government-opposition ideological range, as previous work suggests that parliamentary majorities may be reluctant to extend oversight procedures when oppositions are motivated to obstruct the government (Sieberer et al. 2020). This value is calculated as a seat-weighted average of the distances between each opposition party and the ideal cabinet position. Third, I control for government minority status. Previous work has argued that minority cabinets may have a special motivation to create strong committee procedures (Strøm 1990). The fourth control is the (logged) number of days a cabinet spent in office. Fifth, I control for the time trend as previous work finds an upward trend in committee strength in European democracies (Martin and Vanberg 2020). The time trend is coded using a dummy variable for the decade in which a given cabinet started life. Lastly, I include country fixed effects to account for any unobserved country heterogeneity. See Appendix J for descriptive statistics.

Identification

In relation to the model specified in Equation (5), I note that robust identification in joint likelihood models is facilitated when their specification satisfies the exclusion restric-

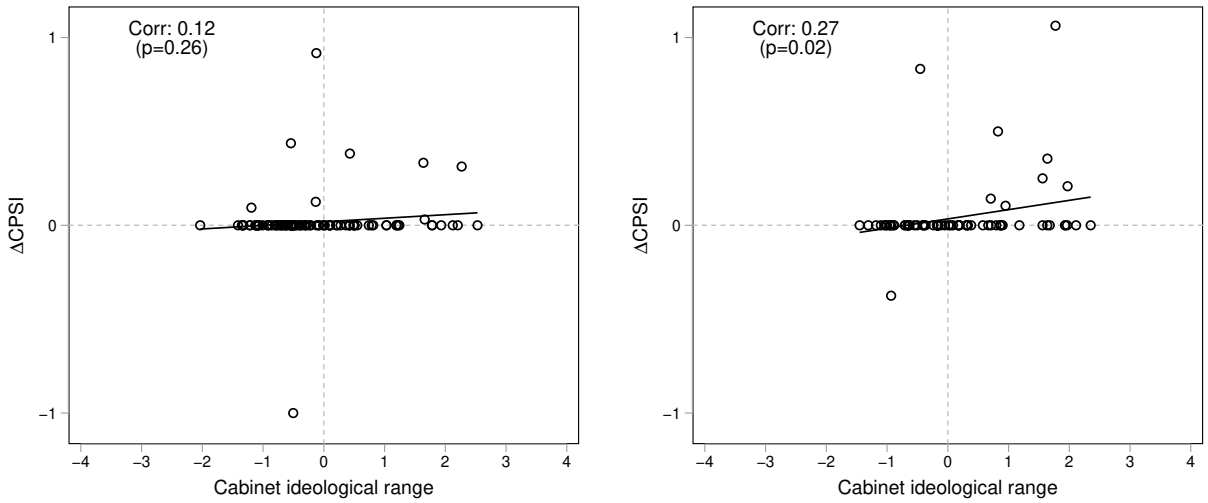
tion requirement, i.e. when there are strong predictors of selection which do not have an independent impact on outcomes (Chiba, Martin, and Stevenson 2015; Chiba and Johnson 2019). In my specification as discussed above many of the variables included in the selection equation do not appear in the outcome equation (and vice versa), and a plausible case can be made that at least some of such covariates have an effect on the implemented procedure only through government selection. In particular, while previous research has shown that the presence in a potential cabinet of the party of the median legislator or the previous PM affects the formation probability, there is little prior reason to think that the inclusion of such parties in the government will affect the implemented committee procedures except through government selection.

EMPIRICAL RESULTS

Bivariate Correlations

I begin by examining the bivariate correlations between the CPSI index values and cabinet ideological divisions given the parties' proposing probability similarity. In my theoretical model higher values of parameter ρ are associated with *more* costly oversight, while higher CPSI values indicate *less* costly oversight. This implies that if my theory has some empirical plausibility, I expect to find a *positive* relationship between coalition ideological divisions and the change in CPSI that parties implement, and this relationship should become stronger as PPS increases. Figure 1 investigates whether this is the case in my sample of 174 multi-party governments. In both panels the x-axis plots country-standardized cabinet ideological range and the y-axis shows the change in CPSI index between the start date and the end date of a government. Panel (a) displays countries with below-average country-standardized PPS,

and panel (b) those with above-average country-standardized PPS. These figures reveal a small positive association between CPSI change and cabinet ideological range which is indeed stronger in the subset of coalitions with above-average PPS values.



(a) Below-average PPS **(b)** Above-average PPS.

Figure 1: Change in CPSI by Cabinet Range and PPS

ML Model Without Selection

Table 1 presents the results from ML models without selection (Equation 4) in which I estimate the effects of cabinet ideological range and proposing probability similarity on the implemented committee procedure, controlling for the inherited status quo procedure and confounders. The standard errors are obtained using a robust sandwich estimator. Model 1 estimates direct effects without the interaction. The results from this model show that, all else equal, cabinet ideological range has a positive effect on committee strength. This gives support to Proposition 1. Model 2 includes an interaction between cabinet range and proposing probability similarity. These results indicate that the effect of cabinet ideological

range increases with proposing probability similarity and that this interaction is significant at $p < 0.01$. This finding is in line with Proposition 2. Overall, these results thus provide initial empirical support to the proposition that coalition parties implement increasingly more extensive oversight procedures the higher is the policy preference divergence between parties and the more similar is their proposing probability.

Table 1: ML Model Results (No Selection)

Implemented CPSI	Model 1	Model 2
Status quo CPSI	0.767***	0.760***
	(10.14)	(10.08)
Ideological divisions in coalition	0.063**	0.007
	(2.49)	(0.35)
Proposing probability similarity (PPS)	0.007	-0.127**
	(0.17)	(-2.32)
Ideological divisions in coalition*PPS		0.134***
		(2.79)
Government-opposition divisions	0.038**	0.039**
	(2.27)	(2.37)
Minority government	0.085**	0.076**
	(2.36)	(2.08)
Cabinet duration	0.012	0.011
	(1.14)	(1.07)
Constant	-0.160*	-0.115
	(-1.72)	(-1.30)
Observations	174	174
AIC	-141.35	-145.29

Note: both models include country and decade dummies, z statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Figure 2 shows how the marginal effect of cabinet range in Model 2 changes with proposing probability similarity when keeping other co-variates at their mean values. This figure

reveals that the marginal effect is close to zero and is not significant when the similarity value is approximately below 0.25. However, the effect becomes positive and significant at higher values of proposing probability similarity and increases in magnitude as the similarity grows. The CPSI implemented by a coalition which allocates proposing probability equally is estimated to increase by 0.14 with each unit increase in cabinet range. Keeping other things equal, such a coalition increases the implemented CPSI by 0.25 when its range increases from one standard deviation below the mean to one standard deviation above the mean, and by 0.74 when its range increases from minimum to maximum. Given that the CPSI scores range from 0.167 to 2.75 in my sample these effects are moderate in size.

I further note that among the controls, having a minority status has a positive and significant effect on committee strength. This result is in line with earlier work on minority cabinets and legislative institutions (Strøm 1990). I also find that cabinet-opposition ideological divisions have a positive and significant effect, a finding which is contrary to expectations derived from previous research (Sieberer et al. 2020). I note that cabinet duration does not affect the strength of the committee oversight procedure. Finally, the decade fixed effects indicate that compared to the baseline decade of the 1940s the mean implemented CPSI tends to increase in later decades in the studied period.[21]

ML Model With Selection

It is possible that the results reported in Table 1, Model 2, may be biased by the non-

21. The decade fixed effects are omitted from Table 1. See Appendix K for full results. I note that the upward trend is consistent with the theoretical notion that coalitions generally want to strength oversight institutions refraining from making scrutiny more costly.

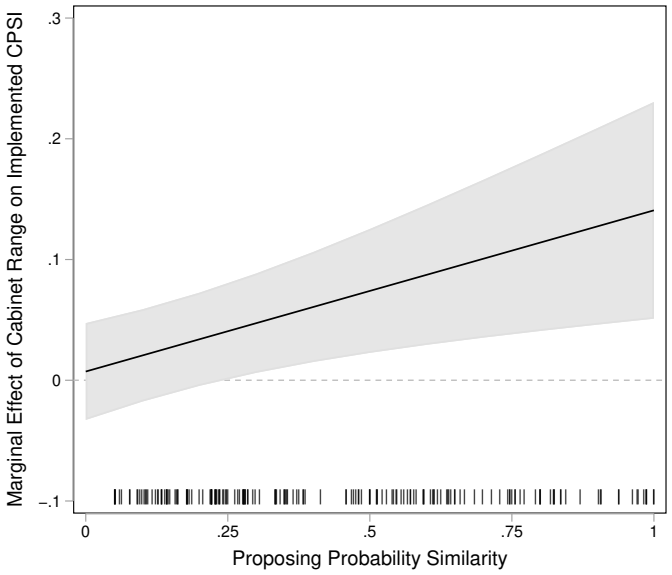


Figure 2: Marginal Effect of Cabinet Ideological Range with 95% CIs

random selection of my sample of coalition governments. This may arise in particular if some unobserved variables affect both the government selection and the choice of procedures. I check for such possible sample selection bias by estimating a joint ML regression model (Equation 5) which uses a Gaussian copula to specify the joint probability for the implemented oversight procedure given the coalition government is selected (Chiba, Martin, and Stevenson 2015). A desirable feature of this joint ML model is that it allows me to estimate the amount of correlation between government formation and procedural choice. This is achieved by estimating the parameter θ which has the permissible range $\theta \in [-1, 1]$ with values below and above zero indicating negative and positive correlation respectively. The parameter θ can be tested for significance, and hence obtaining a θ value which is statistically different from zero would indicate that I should prefer the joint probability model over the univariate model.[22]

22. I follow Chiba, Martin and Stevenson (2015) in estimating and testing the inverse

The joint model results show no selection bias (see Appendix L for details). The estimated correlation parameter is small and positive, but is not statistically different from zero. This finding implies that to the extent that parties choose coalitions based on factors related to expected procedural choices, these factors seem to be adequately accounted for by the observed variables in the equation for the implemented procedure. Importantly, the results for both the univariate and joint models show that multi-party governments implement increasingly more extensive oversight procedures when coalition ideological divisions are high *and* the coalition parties' proposing probabilities are similar. I finally note that the results from the selection part of the joint model are in line with the findings from previous research regarding factors that affect government formation.

Robustness

I perform several robustness tests of my results.[23] First, I probe the sensitivity of the interaction effect in Model 2 to dropping each country one by one. This test demonstrates that the coefficient for the interaction term in such reduced samples remains positive and significant at least at the 0.1 level. I also use Cook's distance to identify cabinets with a significant influence on the model coefficients and re-estimate the models by dropping each in turn. The conclusion is that the interaction effect stays positive and significant at

hyperbolic tangent of the error correlation parameter θ, that is, $\tanh^{-1}(\theta)$.

23. These tests are reported in full in Appendix N. See N1 for jackknifing; N2 for models with two-party coalitions; N3 for results with alternative PPS specifications; N4 for results with portfolio-based PPS measures; N5 for partial CPSI results; N6 for estimation with default RILE; N7 for models with cabinet range difference; N8 for models regressing CPSI change.

least at the 0.05 level. Next, I run the analysis on two-party coalitions. As my theory is derived from a model with two players, I want to address possible concerns about a mismatch between theory and empirical analysis. This test demonstrates that even though my sample of coalition governments is reduced by half, my main results continue to hold with the interaction term staying positive and significant at the 0.1 level. I further investigate whether my main effects differ between established democracies in western Europe and new democracies in eastern Europe by re-estimating Model 1 and 2 in these sub-samples. These results are shown in Table 2. I note that the coefficient for the interaction term in Model 2 is positive and significant ($p < 0.05$) in the analysis of western Europe, but is not statistically different from zero in that of eastern Europe. Interestingly, cabinet range continues to have a direct positive and significant effect in Model 1 in new, but not in established, democracies.

In a further set of robustness tests I re-estimate both models from Table 1 using two alternative specifications of the PPS measure. The first alternative uses simply the legislative size of the largest party. In this analysis I find that my main results remain unchanged. The other alternative uses the size of the largest and the smallest party. Here my results hold only if I focus on minimum winning coalitions, a finding that makes sense given that only in such cabinets smallest parties are critical for bargaining over procedures. In a further robustness check I re-estimate my results by measuring parties' proposing probability similarity based on party shares of cabinet posts. To perform this test I combine my government data with a comprehensive dataset on cabinet membership in European democracies (see Casal Bértoa and Enyedi 2022). These results indicate that while the main effects continue to hold in established democracies, I do not find support for my theory in new democracies. I note that these differential effects echo the findings from the sub-sample analysis reported above.

Table 2: ML Model Results (No Selection): Established vs New Democracies

Implemented CPSI	Established		New	
	Model 1	Model 2	Model 1	Model 2
Status quo CPSI	0.928***	0.914***	0.581***	0.585***
	(12.69)	(12.23)	(5.18)	(5.13)
Ideological divisions in coalition	0.009	-0.019	0.175***	0.223**
	(0.72)	(-1.38)	(3.53)	(2.04)
Proposing probability similarity (PPS)	-0.046*	-0.145**	0.266*	0.318*
	(-1.75)	(-2.24)	(1.73)	(1.91)
Ideological divisions in coalition*PPS		0.093**		-0.087
		(2.14)		(-0.56)
Government-opposition ideological divisions	0.014	0.017	-0.046	-0.048
	(1.07)	(1.30)	(-0.80)	(-0.83)
Minority government	0.048**	0.036*	0.169**	0.170**
	(2.25)	(1.89)	(2.50)	(2.51)
Cabinet duration	0.014	0.012	0.046*	0.046*
	(1.14)	(1.02)	(1.91)	(1.90)
Constant	-0.065	-0.038	-0.514***	-0.544***
	(-0.81)	(-0.49)	(-2.82)	(-2.77)
Observations	109	109	65	65
AIC	-143.61	-144.40	-31.82	-30.06

Note: all models include country and decade dummies, z statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

In the final set of robustness checks, I re-estimate both models using partial versions of the CPSI index: CPSI-P and CPSI-M. This analysis finds that the effects of primary interest are substantively the same regardless of the version of the CPSI index. I further re-estimate Model 1 and Model 2 using ideological range variables computed based on the standard RILE measure of party positions. In this test I find that the main effects remain substantively unchanged. I also estimate Model 1 and Model 2 by introducing an additional control for how much more divided the current coalition government was than the previous coalition. I find that the main effects hold in this re-estimation. Finally, I also run my models by regressing CPSI change rather than the implemented CPSI level on co-variates without controlling for the status quo CPSI. This analysis reveals that the interaction term in Model 2 remains positive and significant at the 0.05 level.

DISCUSSION

The analysis finds empirical patterns which are broadly consistent with my model of procedural choice under coalitions governments. Legislative institutions tend to be strengthened more extensively under multi-party coalitions in which preference divergence between coalition partners is substantial and in which coalition parties have similar probability of making proposals to implement coalition policy. These theoretical and empirical findings have important implications for coalition research. For one thing, my results confirm that the design of political institutions may be driven by strategic choices of parties that form coalition governments in parliamentary democracies. In particular, legislative committee procedures can – at least to some extent – be viewed as self-imposed constraints that coalitions construct to facilitate multi-party governance. Moreover, the results offer an important

refinement to existing research on how coalitions shape legislative institutions in parliamentary democracies. I find that strong committee oversight procedures may not always arise as a straight-forward response to preference divergence between coalition partners. Rather, the effect of such preference divergence may be critically moderated by the extent to which gains from oversight procedures are substantial for *all* coalition parties.[24]

The present work opens up three promising avenues for further work on how coalitions shape legislative institutions in parliamentary democracies. First, the model of procedural bargaining I propose assumes that a party forming a coalition agrees to undertake a procedural reform only if the marginal benefit exceeds the marginal cost for that party. While this may be a reasonable assumption in many contexts, it is also possible the wider bargaining environment can offer parties opportunities to trade concessions across procedures and other benefits. To make progress in this direction, we would need a model of coalition bargaining in which parties negotiate over procedures as well as over other goods. To my knowledge existing coalition research has not yet offered such a model. Theorizing the role of procedures in coalition bargaining would thus be an important next step for the current analysis.

Second, I find a somewhat different institutional dynamic in western and eastern European democracies. The result is consistent with previous studies which show that a country's length of experience with democracy affects procedural choice outcomes (e.g. Carroll, Cox, and Pachón 2006). In the context of my analysis, parties in mature democracies may have more precise expectations than parties in new democracies regarding the extent to which

24. These results are consistent with Wang's (2013) cross-country analysis of committee blocking power.

they will be able to shape the coalition policy agenda. It is also possible that the length of experience with electoral democracy affects the level of certainty that parties face regarding the effect of procedural reforms. In its current form, my model does not yet capture such important nuances which I leave for future research to address systematically.

Third and finally, my analysis suggests that the role of opposition parties in shaping legislative institutions, and committee procedures in particular, may require closer attention. Based on insights from previous work (Sieberer et al. 2020) I expected that since committee scrutiny can be argued to improve the position of non-cabinet parties, high conflict between government and opposition would militate against extensive expansions in the strength of committee procedures. In contrast, I find a positive association between cabinet-opposition ideological range and expansions in committee strength. This suggests that a high conflict with the opposition may in fact create conditions conducive to stronger oversight. To understand this dynamic systematically my model of procedural choice would need to be extended to incorporate non-cabinet actors. I leave this important task for future research.

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